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COMPARISON OF IRT OBSERVED-SCORE AND TRUE-SCORE
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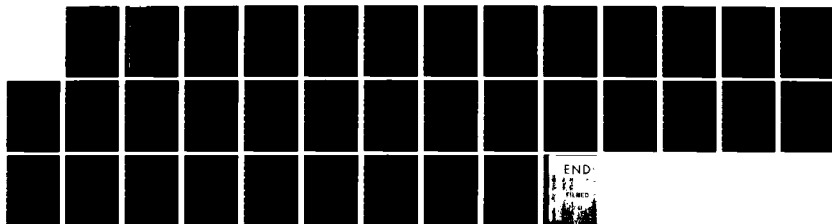
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TRUE-SCORE 'EQUATINGS'

Frederic M. Lord

and

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This research was sponsored in part by the
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Frederic M. Lord, Principal Investigator



Educational Testing Service
Princeton, New Jersey

July 1983

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Abstract

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Comparison of IRT Observed-Score and True-Score 'Equatings'*

Most IRT equating is currently attempted by the true-score equating procedure described in Lord (1980, Chapter 13). Lord also describes an IRT observed-score procedure, which until now seems not to have been further investigated, perhaps because it is more complicated and more expensive than the true-score procedure. The present article reports an empirical research study comparing the results of applying these two procedures to real test data.

Sections 1 and 2 outline the true-score and the observed-score procedures, respectively. Section 3 discusses the theoretical advantages and disadvantages of each procedure. Section 4 describes the real test data used to provide a comparison of the two methods. Section 5 describes the procedures for estimating item and ability parameters. Section 6 reports and summarizes the empirical results.

Item response theory models the probability of a correct response by an examinee to a test item as a monotonically increasing function of ability. The model used here is Birnbaum's three-parameter logistic model given by the following formula:

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$$P_i(\theta_a) = c_i + (1 - c_i)/(1 + \exp(-1.7a_i(\theta_a - b_i))) \quad (1)$$

where $P_i(\theta_a)$ is the probability of examinee a getting item i correct

b_i is the difficulty of item i ;

a_i is the discrimination index for item i ;

c_i is the lower asymptote for item i ;

θ_a is the ability of examinee a ($-\infty < \theta_a < \infty$).

$P_i(\theta_a)$ has a minimum of c_i and a maximum of 1. This model assumes that the test is unidimensional.

1. True-Score Equating

Since the expected score of examinee a on item i is $P_i(\theta_a)$, the examinee's expected number of right answers is $\sum_i P_i(\theta_a)$. In classical test theory, this expectation is called the (number-right) true score, $\xi_a \equiv \sum_i P_i(\theta_a)$. For the moment, we do not deal with the scores of particular examinees, so the subscript a will be dropped. Here the true score for test X containing n items is the mathematical variable

$$\xi = \sum_{i=1}^n P_i(\theta) \quad , \quad (2)$$

a monotonic increasing function of θ . If test Y contains m items and measures the same ability θ as test X , the true score on test Y is the mathematical variable

$$\eta \equiv \sum_{j=1}^m P_j(\theta) \quad . \quad (3)$$

The variables ξ , η , θ are all measures of the same psychological trait, they differ only in the numerical scale on which the measurements are expressed. Thus true scores $\xi = \xi_0$ and $\eta = \eta_0$ corresponding to any given $\theta = \theta_0$ represent identical levels of ability. Any examinee whose true score on test X is ξ_0 must automatically have a true score on test Y of exactly η_0 , provided the IRT model holds. The situation is the same as when we say that 32° Fahrenheit has the same meaning as 0° Celsius, except that these temperature scales have a linear relationship, whereas the true-score scales have a nonlinear relationship. Thus, ξ_0 and η_0 are equated true scores; this is true in a much stronger sense than is usually implied by the term equated.

In IRT true-score equating, estimated item parameters are substituted into (2) and (3) and a table of corresponding values of ξ and η is calculated. This constitutes the true-score equating table. This table is then applied in practice as if the true scores

were observed number-right scores. Since observed scores have different properties than true scores, this last step has no clear theoretical justification. It is done as a practical procedure, to be justified only by whatever usefulness and reasonableness can be empirically demonstrated for the results.

2. IRT Observed-Score Equating

If the assumptions of IRT hold (as is assumed throughout), the probability that an examinee of ability θ will have a number-right score of $x = 1$ on a two-item test is $P_1 Q_2 + Q_1 P_2$, where $P_1 \equiv P_1(\theta)$ and $Q_1 \equiv 1 - P_1$. The probability that this examinee's score is 0 or 1 is $Q_1 Q_2$ or $P_1 P_2$ respectively. These probabilities constitute the conditional frequency distribution $f_2(x|\theta)$.

If a third item is added to this test, the distribution of x is now

$$f_3(x|\theta) = Q_3 f_2(x|\theta) + P_3 f_2(x - 1|\theta) \quad (x = 0, 1, \dots, 3).$$

where $f_r(x|\theta) = 0$ if $x < 0$ or $x > r$. Using this recursive procedure, a computer can readily determine $f_r(x|\theta)$, even for an n of several hundred.

If the θ of each examinee is known, the (marginal) distribution of x for a group of N examinees is

$$\frac{1}{N} \sum_{a=1}^N f_n(x|\theta_a) \quad . \quad (4)$$

If an m -item test Y yields number-right score y and measures the same ability as test X , then the (marginal) distribution of y for a group of M examinees is

$$\frac{1}{M} \sum_{b=1}^M f_m(y|\theta_b) \quad . \quad (5)$$

A monotonic transformation of the y scores can now be found from (4) and (5) such that the distribution of the transformed y scores is the same as the distribution of the (untransformed) x scores, except for irregularities due to the fact that x and y can only assume integer values. This is done by finding, for each y score, the x that has the same percentile rank in (4) that y has in (5). The x so found is the desired transformed y score.

If the examinees who took test Y have the same distribution of θ as the examinees who took test X , then the resulting transformation of y is an 'equipercentile equating' of the y scale to the x scale. Within groups similar to the groups used to derive the transformation, it has the valuable property that if a cutting score

is chosen on the x scale and the same cutting score is used on the transformed y scale, the proportion of test X examinees selected will be the same as the proportion of test Y examinees selected. This property is essential if test X and test Y examinees are both to be treated equitably, so that an examinee cannot complain that he was injured by the choice of test administered.

When the groups taking tests X and Y are known to have approximately the same distribution of θ (for example, they are two random samples from the same population), there is no reason to use IRT. It is much simpler to do the equipercentile equating using the actual sample distributions of x and y , instead of (4) and (5). The need for IRT arises when the ability distributions of the two groups may differ. In this case, IRT may allow us to estimate the (marginal) frequency distributions of number-right scores that would have resulted if all examinees had taken both tests, without practice or fatigue effects.

In order to do this, the item and ability parameters in (4) and (5) must all be on the same scale. This is usually accomplished by administering a suitable 'anchor test' to both groups of examinees. All answer-sheet responses for both groups are used in a single computer run that estimates all parameters on the same scale. These estimates are then used in (4) and (5), substituting $N + M$ for N or M , to obtain the distributions of x and y for the combined group of $N + M$ examinees. Equipercentile equating of y to x is then carried out in the usual way.

3. Theoretical Perspectives

Practical workers, with the need for equating scores on two different test forms, have over the years used widely different methods (see Angoff, 1971) in an attempt to approximate the desired result. Each practical worker, needing a word to describe his results, asserts that he has produced an equating of y to x . Yet different methods and different groups do not produce identical 'equatings'.

Braun and Holland (1982, page 14) state: "There is some disagreement over what test equating is and the proper method for doing it." They then adopt the definition "Form-X and Form-Y are equated on [population] P " if the distribution of the transformed y scores in population P is the same as the distribution of the (untransformed) x scores.

This definition of the phrase 'equated on population P ' is beyond reproach. One problem, however, is that the qualifying phrase 'on population P ' is typically dropped by the practical worker who writes a research report or publishes an equating table in a test manual.

Unfortunately (as will be shown later in this section) two tests that are equated on population P will typically not be equated for various subpopulations that are included in P . Test scores that are equated for the population of college applicants may well be equated neither for the population of female college applicants, nor for the

the population of male college applicants. The scores are still less likely to be equated for a subpopulation characterized by interest in science, or in music. For the subpopulation of Harvard applicants, the situation is much worse.

If the proportion of applicants admitted to Harvard differs significantly depending on whether they were given form X or form Y of the test, it is clear that the 'equating' was unsuccessful. Since similar inequalities are likely to characterize any equating on any specified population, it may be best not to say that the tests are 'equated' at all, or to simply say that they are 'approximately equated.'

From a practical point of view, the approximation may be quite satisfactory for many subgroups. It is unlikely, however, that the equating will be adequate for any subpopulation having a mean and variance of ability that is sharply different from the mean and variance of the total population used to derive the equating transformation. Extensive practical data illustrating the adequacies and the inadequacies of approximate equatings are given in the 30-volume Anchor Test Study (Loret, Seder, Bianchini, and Vale, 1974).

For a theoretical discussion of alternative equating methods, however, it is important not to start out with a definition of equating that is clearly inadequate for subpopulations of examinees. Given that the IRT model holds, IRT observed-score equating would, for example, be automatically endorsed by the Braun and Holland definition, since their definition mandates equipercentile equating. IRT true-score equating would be definitely rejected by their definition, since in general it

will not lead to x scores and transformed y scores having the same frequency distribution, unless X and Y are strictly parallel forms that are identical in difficulty, in reliability, and also in most other respects.

The important virtue of IRT true-score equating is that if the IRT model holds, the true scores are clearly equated for all subpopulations of examinees. This results from the invariance of IRT parameters across populations of examinees, assumed by the IRT model. The clear flaw in IRT true-score equating is that it equates true scores, not the actually observed fallible scores. Treating observed scores as if they were true scores cannot be justified on any theoretical grounds.

The virtue of IRT observed-score equating is that in a group like that used to derive the equating, any cutting score will accept the same percentage of examinees regardless of the test administered. The flaw is that this holds only for that total group and not for other groups or subgroups.

This last statement is most clearly seen from a very extreme example. Suppose forms X and Y have the same number of items, measure the same ability θ , but differ in difficulty. If the equipercentile equating is carried out on a group of examinees all of whom are guessing at random on almost all the items, the difference in difficulty between the two forms will not manifest itself and any equipercentile equating will approximate an identity transformation of score y . If a slightly more competent group of examinees is used for the equipercentile equating, however, the difference in difficulty

between forms will begin to become apparent and most y scores will be adjusted upwards or downwards accordingly. As the competence of the group used becomes higher and higher, the equating transformation found will differ more and more from the identity transformation found from the original extreme group.

As a second example of the inescapable invalidity of observed-score equating, suppose that tests X and Y are of equal difficulty and that the true scores ξ and η have equal variance, but that y is much less reliable than x . Consider a subgroup of very talented examinees; to make the illustration clear, consider that in this subgroup all examinees have nearly identical θ values. Most of the variation in observed scores x and y is now due to errors of measurement. The equipercentile equating transformation found will thus approximate a straight line with slope

$$\frac{\text{standard deviation of the errors of measurement in } x}{\text{standard deviation of the errors of measurement in } y}.$$

Since y is much less reliable than x , the slope will be much less than 1.

If, on the other hand, the equipercentile equating transformation is found from a group where the true-score variance is large compared to the errors variances, the transformation will tend to approximate a straight line with slope

$$\frac{\text{standard deviation of true scores on } x}{\text{standard deviation of true scores on } y}.$$

Intermediate situations will provide transformations with intermediate slopes. If the wrong equating is applied to any given subpopulation, the population of examinees in the subpopulation accepted will depend on whether they took test X or test Y, an inequitable result.

Our theoretical position, then, is that each method described in Section 2 (as well as all other available equating methods) has its own inadequacies. Since, in practice, some (approximate) equating method must be used, it will be informative to investigate empirically how the two methods of Section 2 compare in a specially contrived practical situation where the correct equating is actually known in advance.

4. Data

These two equating methods were used to equate the chain of six SAT verbal tests described by Petersen, Cook and Stocking in the report IRT Versus Conventional Equating Methods: A Comparative Study of Scale Stability. The tests in this chain were selected such that the first test and the last test are the same. Each test is equated to the next test in the chain using an anchor test. Figure 1 is a diagram of the chain. The capital letters represent the test form, the small letters represent the anchor test. Scores on form V4 are equated to scores on form X2 using the anchor test fe. These equated scores on X2 are equated to scores on form Y3 using the anchor test fm. This gives us an equating of form V4 to Y3. In this manner, one proceeds through the chain, with the final equating of Z5 to V4 giving us a

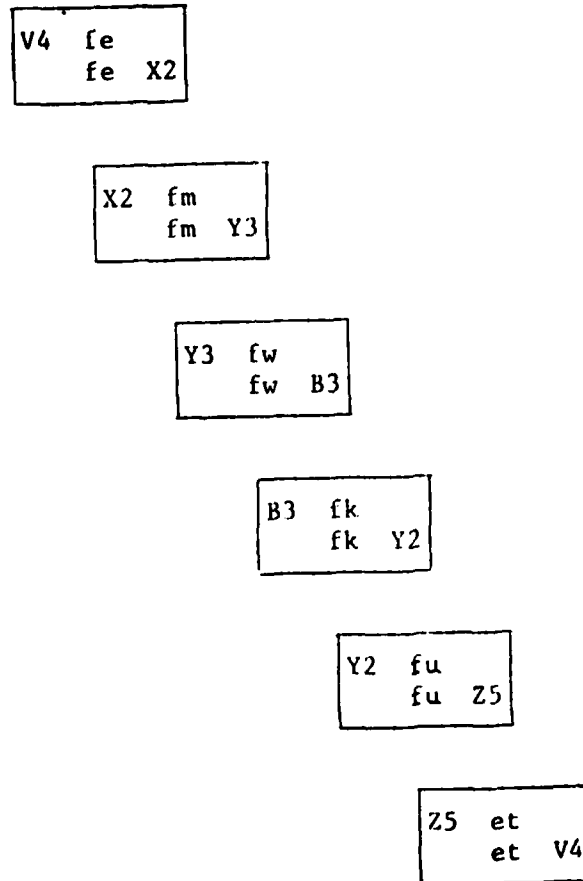


Figure 1. Chain of six SAT verbal equatings. Upper case letters designate test forms; lower case letters designate anchor tests.

table of scores on the original V4 equated to the scores on the V4 at the end of the chain. Any deviation from equality between the two sets of scores could be attributable to scale drift or lack of model fit.

Each form in the chain has 85 items except form V4 which has 90 items. Each anchor test has 40 items. For each form there are two samples of examinees; each sample taking a different anchor test. The two groups taking each form were random samples from the same population for all of the forms except Y3. For the parameter estimation runs a random sample of approximately 2670 examinees was selected from the data obtained at the test administration of that form and anchor test.

5. Parameter Calibration

The item parameters and abilities were estimated by a modified version of the computer program LOGIST, (Wood, Wingersky, & Lord, 1976) in six separate calibration runs. In Figure 1, each box (containing two forms and one anchor test) represents one LOGIST run. The item responses for items not taken by an examinee, such as the X2 items for examinees taking form V4 in box 1, are treated as not reached items.

All of the estimated parameters within each LOGIST run are on the same scale and either method of equating can be used to equate the scores for the two tests. The anchor tests are not used directly in the equating but are used in LOGIST so that the estimated parameters within a LOGIST run are on the same scale.

6. Results

In using the IRT observed-score equating method, two estimated distributions of observed scores are equated so that the transformed y scores and the (untransformed) x scores have the same distribution. Figure 2 is presented to demonstrate, that at least for one set of data, this estimated distribution of observed scores is a reasonable fit to the actual distribution of observed scores. The frequencies are plotted against formula scores which are the number right minus a fraction of the number wrong. The fraction is one over the number of choices. Since the estimated observed-score distribution can only be obtained for number-right scores, the transformation to formula scores assumes that there are no omits, that is, that the number wrong is the total number of items minus the number right. In order to compare the two distributions, the observed-score distribution should be based on a group that has no omits. Consequently, a form of the SAT verbal different from the ones in the chain was used for this Figure in order to get a sufficiently large enough sample for the frequency distribution and for the item calibration.

The agreement shown in Figure 2 is good except that the tails of the estimated distribution are too high. This discrepancy is presumably due to the use of estimated θ in place of true θ for the practical implementation of (4). Since a similar discrepancy affects

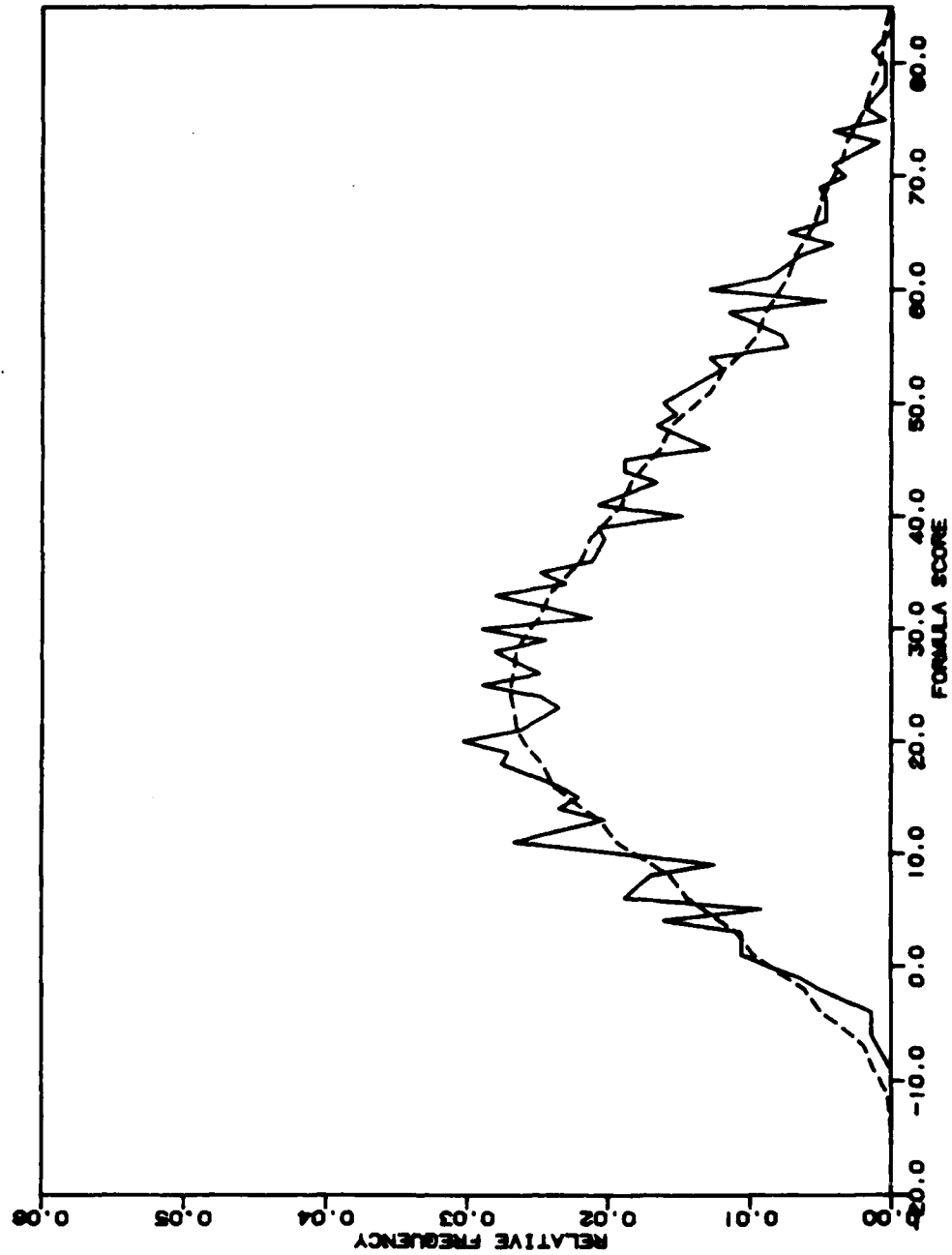


FIGURE 2
COMPARISON OF DISTRIBUTION OF OBSERVED
SCORES AND ESTIMATED DISTRIBUTION OF
OBSERVED SCORES - SAT VERBAL - NO OMTS
--- ESTIMATED OBSERVED-SCORE DISTRIBUTION
— OBSERVED-SCORE DISTRIBUTION

the estimated observed-score distributions of both test X and test Y, the effects of the discrepancies tend to cancel out in the equating process.

In our chain-equating study, each method of equating was applied separately to the whole chain of equatings, resulting in a line for each method equating form V4 at the beginning of the chain to form V4 at the end of the chain. These two lines are plotted in Figure 3 along with a 45° line. The solid line is the IRT true-score equating line; the dotted line, falling practically on top of the solid line, is the IRT observed-score equating line. To equate scores below chance level, that is $\sum_{i=1}^n c_i$, for the IRT true-score line, the method given on pages 210-211 of Lord (1980) was used. For scores above 0, the maximum difference between the two equatings was .2; for scores below 0, the maximum difference was .8 which occurred at the chance level. If the equating methods were perfect and there were no scale drift, the equating line would be the dashed 45° line.

Figure 4 shows the two equating methods applied to one individual link in the chain. This particular link was selected because the IRT true-score equating line between these two forms had the greatest discontinuity in the slope at the chance level. The largest difference between the two lines occurred at the chance level and was 1.6. For scores above 0 the maximum difference between the two lines was .4.

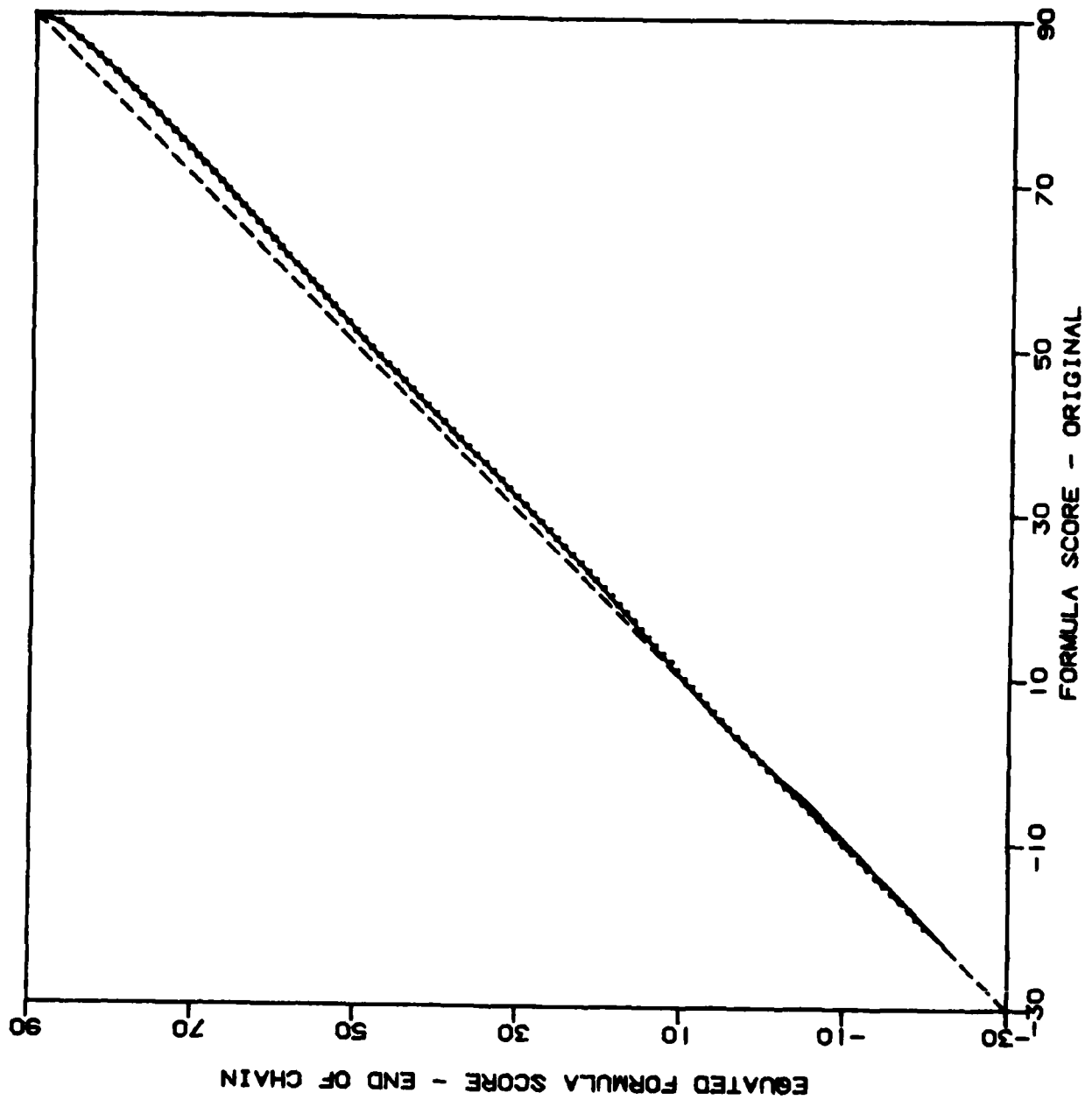


FIGURE 3
COMPARISON OF TRUE-SCORE EQUATING
AND ESTIMATED OBSERVED-SCORE EQUATING
OVER CHAIN OF SIX EQUATINGS
--- ESTIMATED OBSERVED-SCORE EQUATING
--- TRUE-SCORE EQUATING
--- CRITERION

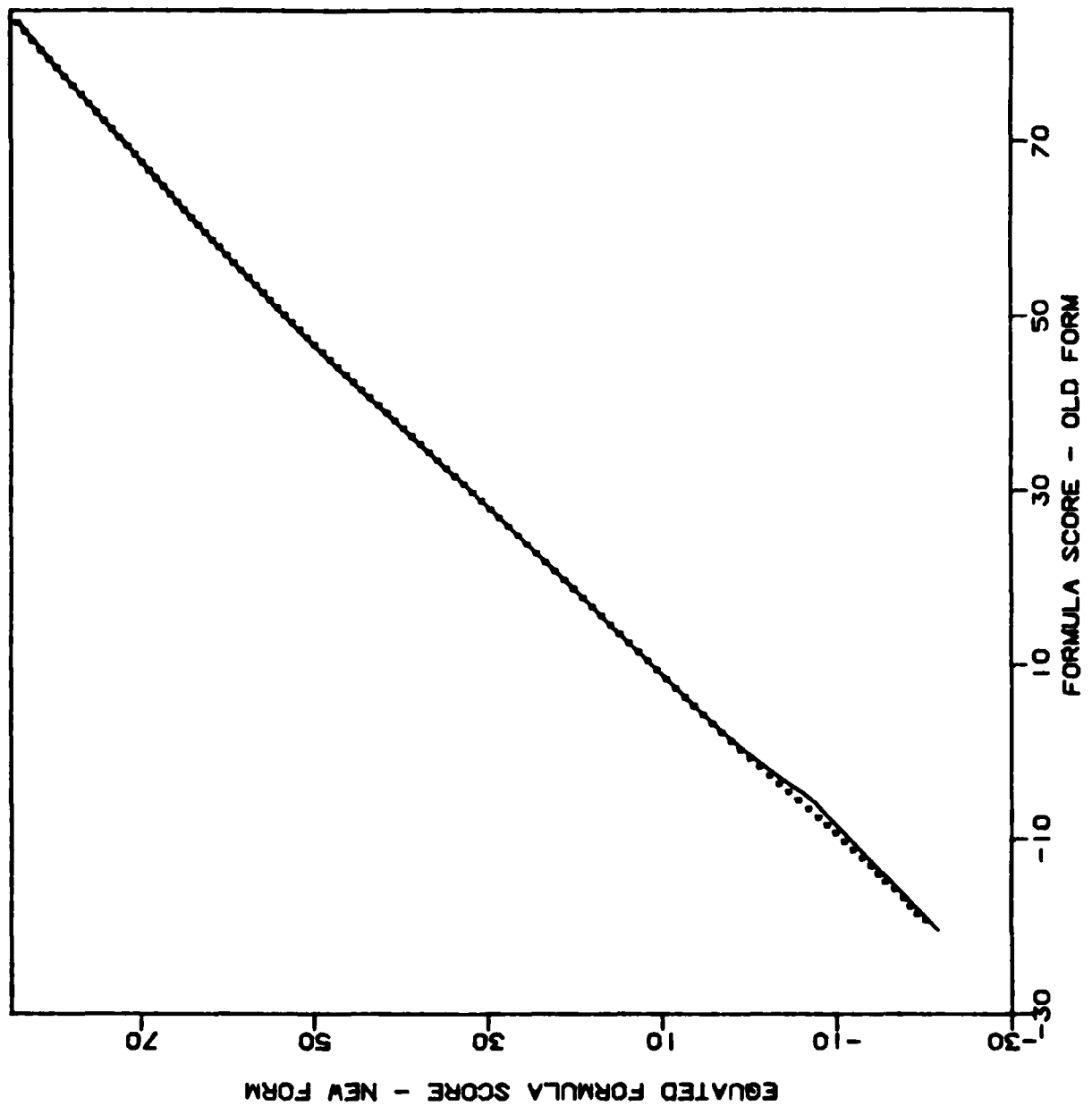


FIGURE 4
COMPARISON OF TRUE-SCORE EQUATING
AND ESTIMATED OBSERVED-SCORE EQUATING
SINGLE LINK IN CHAIN
..... ESTIMATED OBSERVED-SCORE EQUATING
—— TRUE-SCORE EQUATING

Given that there is no clear theoretical justification for applying IRT true-score equating to observed scores and that the equipercentile equating of the IRT observed-score distributions is population dependent, the close agreement between the two lines is reassuring.

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FIGURE 4
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